International Risk Sharing in the Short and Long Run under Country Heterogeneity

*P. Fuleky, **L.Ventura , ***Q. Zhao

*UHERO and Department of Economics, University of Hawaii at Manoa

**Department of Economics and Law, Sapienza, University of Rome

*** Department of Economics, University of Hawaii at Manoa

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Abstract

Empirical tests of international risk sharing usually focus on the short run and impose homogeneity across countries. We extend the existing literature by analyzing both the short and the long run aspects of risk sharing, and by using novel econometric tools that properly account for heterogeneity across a large set of countries. The results confirm that international risk sharing is only partial and quite variable across economies. However, there appears to be no trade-off between the short and long run: countries better insured in the long run also tend to perform better in the short run, with the strength of domestic credit markets being crucial in shaping this relationship.

Keywords: Panel data, Cross-sectional dependence, Heterogeneous effects, International risk sharing, Consumption insurance

JEL codes: C23, C51, E21, F36

1 Introduction

The pioneering work of Cochrane (1991), Mace (1991), and Obstfeld (1994), set the path for a number of consumption risk sharing tests presented in the recent literature. The empirical implementation of these tests tends to rely on a crucial assumption of homogeneity across countries, which is unlikely to hold in large, worldwide panels (Obstfeld, 1989, 1994). Specifically, all economies are assumed to be characterized by similar time and risk preferences, uniform dimensions, and affected by global shocks to the same degree. Using these homogeneity assumptions, a measure of risk sharing is usually estimated by a fixed effects panel data regression of cross-sectionally demeaned consumption on cross-sectionally demeaned income (see for example Kose et al., 2009; Sorensen et al., 2007; Sorensen and Yosha, 2000). Under imperfect insurance, the effect of income shocks on consumption indicates the relative shortfall of risk sharing.

If the homogeneity assumptions underlying the analysis are violated, the estimation results may be biased. Schulhofer-Wohl (2011) shows that heterogeneity, if not properly accounted for, may lead to incorrect conclusions about the degree of risk sharing. We extend the existing literature to allow for variation in preferences and exposure to aggregate risk across economies. Our methodology is specifically designed to deal with cross-sectional dependence and heterogeneity in a wide panel of countries. In principle, greater heterogeneity promotes higher levels of risk sharing (see Kalemli-Ozcan et al., 2005, 2001). For example, asset trading can foster risk sharing if domestic and foreign asset returns are asymmetric. The benefit will be limited if, due to geographic, political or cultural proximity, countries only engage in asset trading with partners that are affected by similar shocks. On the other hand, asset trading between countries that experience contrasting shocks is expected to result in a greater degree of insurance. However, even our robust econometric tools are unable to overturn the main conclusions of previous studies on this topic as we find little evidence in support of full international risk pooling.[1] Nonetheless, we observe sizable variation in risk sharing across the world, and for some countries the extent of risk sharing appears to be different from previous empirical findings.

In addition to accounting for cross-sectional heterogeneity, we also take into consideration the presence of common stochastic trends in the data, and analyze risk sharing in an error correction framework (see Davidson et al., 1978). This allows us to contribute to the discussion about the nature of risk sharing, particularly whether it should be considered a long or a short run phenomenon. In principle, risk sharing at these horizons might display contrasting patterns due to unequal availability of smoothing channels in the long and short run. On the one hand, Leibrecht and Scharler (2008) and Artis and Hoffmann (2012) argue that that the scarcity of international risk sharing documented in the literature is mainly due to a lack of insurance against permanent shocks. On the other hand, Pierucci and Ventura (2010) find that the sensitivity of consumption to idiosyncratic shocks is similar in the short and long run.

Whether risk sharing at one horizon exceeds that at another, depends on the path of adjustment towards equilibrium between income and consumption. If an idiosyncratic income shock sets in motion incremental adjustments in consumption, its immediate impact will be smaller than its long run impact. In this case short run risk sharing will appear to be greater. However, if consumption instantaneously overshoots the equilibrium, the immediate impact of the shock will exceed the long run effect. In this case long run risk sharing will appear to be greater. Our results indicate that the former scenario is the typical one, and that the extent of risk sharing tends to be somewhat greater in the short run. In addition, we observe a positive relationship

[1]The lack of international consumption risk sharing was recognized in the pioneering work of Asdrubali et al. (1996); Lewis (1996, 1997); Obstfeld (1994), and was corroborated recently by Bai and Zhang (2012); Callen et al. (2011); Corsetti et al. (2008).

between risk sharing at the two horizons: countries characterized by less consumption smoothing in the long run are also those heavily affected by idiosyncratic shocks in the short run.

Our contribution to the international risk sharing literature has several important components. First, we re-evaluate earlier results indicating an absence of perfect risk sharing using a more flexible econometric model that allows for general heterogeneity across countries. Second, we embed the analysis in a dynamic framework and show that long term and short term risk sharing are inextricably linked. Specifically, we compare risk sharing in the short and long run while controlling for country heterogeneity within an error correction model. Third, we use a larger panel containing data on 158 countries and exhibiting greater heterogeneity than earlier studies did. While the existing literature has focused on smaller, more homogeneous sets of countries, our sample allows us to analyze risk sharing in previously overlooked regions. Fourth, we explore how various macroeconomic factors affect international risk sharing, and find that the depth of domestic credit and financial markets plays an important role in addition to the general level of development.

The rest of the paper is organized as follows: Section 2 contrasts the conventional and the proposed methodologies to estimate the extent of risk sharing, Section 3 presents our empirical analysis, and Section 4 concludes.

2 Methodology

Under perfect risk sharing, utility maximization ensures that marginal utility growth is equalized across countries and depends on aggregate consumption (or aggregate income) but not on idiosyncratic shocks (Cochrane, 1991; Mace, 1991; Obstfeld, 1994). If preferences are represented by CRRA utility functions, the risk sharing hypothesis can be tested using the equation

$$c_{it} = \alpha_i + \gamma_i^c \bar{c}_t + \beta_i x_{it} + \varepsilon_{it} \ , \ \ i = 1 \dots N, \ \ t = 1 \dots T \ , \tag{1}$$

where c_{it} is a measure of consumption for country i, $\bar{c}_t$ is a measure for aggregate consumption, and x_{it} is an idiosyncratic variable. Full risk sharing implies $\beta_i = 0$ and $\gamma_i^c > 0$, and under the double assumption of homogeneous discount factors and coefficients of relative risk aversion $\gamma_i^c = 1$. However, Obstfeld (1989) found little empirical support for the hypothesis $\gamma_i^c = 1$ even in countries with similar characteristics, such as Germany, Japan and the United States. Still, many papers in the field (including Kose et al., 2009; Sorensen et al., 2007; Sorensen and Yosha, 2000) have built on these homogeneity assumptions. Restricting the analysis to a pool of reasonably similar countries, they use the simplified equation

$$c_{it} - \bar{c}_t = \alpha_i + \beta_i x_{it} + \varepsilon_{it} \ . \tag{2}$$

to test the null hypothesis $H_0 : \beta_i = 0$, where β_i is commonly interpreted as the extent of the departure from perfect risk sharing. The variable x_{it} typically represents a proxy for idiosyncratic income, computed as the difference between the particular country's income and a measure of aggregate income, $y_{it} - \bar{y}_t$. The aggregates, $\bar{c}_t$ and $\bar{y}_t$, are conventionally estimated by cross-sectional averages of consumption and income, respectively. To obtain an overall β coefficient for a set of countries, most researchers pool the data and estimate the fixed effects regression

$$c_{it} - \bar{c}_t = \alpha_i + \beta(y_{it} - \bar{y}_t) + \varepsilon_{it} \ , \tag{3}$$

which introduces an additional level of homogeneity.

Our analysis departs from the conventional methodology. We rely on the framework behind the common correlated effects (CCE) estimator of Pesaran (2006) to account for heterogeneity among countries. The CCE estimator uses cross-sectional averages to filter out common factors from linear relationships among heterogeneous panels. It is equivalent to ordinary least squares estimation of the β_i coefficient in a regression augmented by the cross-sectional means of the variables

$$c_{it} = \alpha_i + \beta_i y_{it} + \gamma_i^c \bar{c}_t + \gamma_i^y \bar{y}_t + \varepsilon_{it} \ , \tag{4}$$

which is the reduced form of

$$c_{it} - \gamma_i^c \bar{c}_t = \alpha_i + \beta_i (y_{it} - \tilde{\gamma}_i^y \bar{y}_t) + \varepsilon_{it} \ , \tag{5}$$

The country specific $\gamma_i^y = -\beta_i \tilde{\gamma}_i^y$ and γ_i^c coefficients allow the amount of income and consumption driven by global shocks to vary across economies. In particular, they robustify the estimation against heterogeneous preferences and variation in the extent of global shocks transmitted to individual countries. Further merits of this methodology relative to the conventional one are discussed in depth by Fuleky et al. (2014).

Although most empirical analyses to date tested the risk sharing hypothesis with differenced data, several recent studies, including Becker and Hoffmann (2006) and Artis and Hoffmann (2012), provided a rationale for risk sharing in the long run and argued in favor of estimating risk sharing equations in levels.[2] If c_{it} and y_{it} contain stochastic trends, but the error term of equation (4) is void of unit roots, then the relationship between the variables can also be estimated within an error correction model (Leibrecht and Scharler, 2008; Pierucci and Ventura, 2010). For an individual

[2]Taking a different approach, Flood et al. (2009) advocated the analysis of risk sharing using the variance of unsmoothed consumption.

country, the deviation from the long run equilibrium between idiosyncratic income and consumption is captured by the residual of model (4) estimated with data in levels. The speed at which this equilibrium error is corrected, κ_i, can then be estimated along with the immediate impact of income changes, β_i^{SR}, in the error-correction model

$$\Delta c_{it} = \alpha_i^{SR} + \kappa_i \hat{\varepsilon}_{it-1}^{LR} + \beta_i^{SR} \Delta y_{it} + \gamma_i^{c,SR} \overline{\Delta c}_t + \gamma_i^{y,SR} \overline{\Delta y}_t + \varepsilon_{it}^{SR} \ , \tag{6}$$

where $\hat{\varepsilon}_{it}^{LR} = c_{it} - \hat{\alpha}_i^{LR} - \hat{\beta}_i^{LR} y_{it} - \hat{\gamma}_i^{c,LR} \bar{c}_t - \hat{\gamma}_i^{y,LR} \bar{y}_t$ is the long run equilibrium error after controlling for permanent global shocks. In equation (6), the heterogeneous impact of transitory global fluctuations is filtered out via country specific coefficients, $\gamma_i^{c,SR}$ and $\gamma_i^{y,SR}$, assigned to cross-sectional means of differenced consumption and income, $\overline{\Delta c}_t$ and $\overline{\Delta y}_t$, respectively.

In addition to the direct effect of an income shock, disequilibrium in the system will also prompt a change in consumption. Whether the equilibrium-error correction adds to or deducts from the direct effect depends on the relative signs of Δy_{it} and $\hat{\varepsilon}_{it-1}^{LR}$. While β_i^{LR} is associated with the level of risk sharing in steady state, the reversal to equilibrium contributes to the short run dynamics, and therefore complements β_i^{SR} or the level of risk sharing in the short run. Equation (6) can also be obtained by a reparameterization of an augmented autoregressive distributed lag, or ADL(1,1), model of the form

$$c_{it} = \alpha_i + \beta_{1i} c_{it-1} + \beta_{2i} y_{it} + \beta_{3i} y_{it-1} + \gamma_{1i}^c \bar{c}_t + \gamma_{2i}^c \bar{c}_{t-1} + \gamma_{1i}^y \bar{y}_t + \gamma_{2i}^y \bar{y}_{t-1} + \varepsilon_{it} \ . \tag{7}$$

The correspondence among the coefficients of models (6) and (7), $\beta_i^{SR} = \beta_{2i}$, $\beta_i^{LR} = (\beta_{2i} + \beta_{3i})/(1 - \beta_{1i})$, $\kappa_i = (\beta_{1i} - 1)$ (see Davidson et al., 1978) implies that $\hat{\beta}_i^{SR}$ will be smaller than $\hat{\beta}_i^{LR}$ whenever $\beta_{2i} > -\beta_{3i}/\beta_{1i}$. The ratio on the right hand side of the inequality captures the relative impact of the two lagged variables. We also explicitly analyze the relationship between short and long run risk sharing using regressions of the form $\hat{\beta}_i^{SR} = a + b\hat{\beta}_i^{LR} + \varepsilon_{it}$. Finally, the cross-sectional dimension of the data panel also enables us to explore how various macroeconomic factors affect international risk sharing. We examine the relationship between the estimated β coefficients and macroeconomic variables using regressions of the form

$$\hat{\beta}_i^{SR} = a^{SR} + b_1^{SR} x_{1i} + b_2^{SR} x_{2i} + \ldots + \varepsilon_{it} \ , \quad \hat{\beta}_i^{LR} = a^{LR} + b_1^{LR} x_{1i} + b_2^{LR} x_{2i} + \ldots + \varepsilon_{it} \ , \quad (8)$$

where the $x_{\cdot i}$ regressors are time averages of country specific macroeconomic indicators.

Having described how our methodology differs from the conventional one, we now turn to our empirical analysis and results.

3 Data and Results

Our study is based on annual data between 1970 - 2010 for 158 countries (Penn World Tables, version 7.1, Heston et al., 2012). With the existing literature largely focusing on smaller sets of rather homogeneous countries, the analysis of such a large heterogeneous panel is a distinguishing feature of our study. To make comparisons across countries and time feasible, we use purchasing power parity converted GDP per capita and consumption per capita at 2005 constant prices. To eliminate the exponential growth pattern in the series, we log-transform the data.

The diagnostic statistics in Table 1 indicate that the log-transformed consumption and income levels are cross-sectionally dependent and non-stationary. The log-differenced series are also cross-sectionally dependent, but they do not contain unit roots. The tests applied to the residuals of equation (4) indicate that the model successfully controls for global shocks and the levels are cointegrated. Therefore the error correction model (6) is an appropriate tool for analyzing short run behavior.

Table 1: Diagnostic Tests

	Levels		Differences		Residuals	
	$\log C$	$\log Y$	$\Delta \log C$	$\Delta \log Y$	Eq.(4)	Eq.(6)
CD	253.45*	239.37*	22.93*	43.33*	0.06	-0.83
$CIPS$	-0.32	1.87	-9.72*	-7.78*	-5.87*	—

Note: Pesaran's (2004) cross-sectional independence test statistic (CD) follows a standard normal distribtion. The 5 % critical value for Pesaran's (2007) panel unit root test ($CIPS$) is -2.06. The lag length for the $CIPS$ test is set to $T^{1/3} \approx 4$. Statistical significance at the 5% level or lower is denoted by *.

Figure 1 illustrates the distribution of individual β_i's. It is immediately obvious that there is a remarkable heterogeneity among countries in terms of their participation in risk-sharing. In line with earlier studies, our estimation results indicate that consumption tends to be affected by idiosyncratic risks in both the long run and the short run ($\hat{\beta}_i \neq 0$), but the extent of risk-sharing tends to be higher in the short run ($\bar{\beta}^{SR} < \bar{\beta}^{LR}$). As explained below, this inequality is implied by a gradual adjustment of consumption to income shocks.

Tables 2 and 3 list the country specific β_i estimates along with the corresponding speed-of-adjustment estimates, κ_i. Our short run results show that 89% of β_i^{SR} estimates are significantly different from zero, and that 26 countries engage in de-smoothing behavior ($\hat{\beta}_i > 1$). The distribution of the long run estimates is shifted slightly higher, with 93% of β_i^{LR} estimates significantly different from zero, and 43 countries engaging in dis-smoothing behavior. From the estimated coefficients we can calculate the mean lag of adjustment towards equilibrium, $\hat{\mu}_i = (1 - \frac{\hat{\beta}_i^{SR}}{\hat{\beta}_i^{LR}})/(-\hat{\kappa}_i)$ (Hendry, 1995). For example, among OECD countries the impact of an idiosyncratic income shock on consumption has an average lag of about half a year, as opposed to about two years and no lag according to the non-CCE-based results of Leibrecht and Scharler (2008)

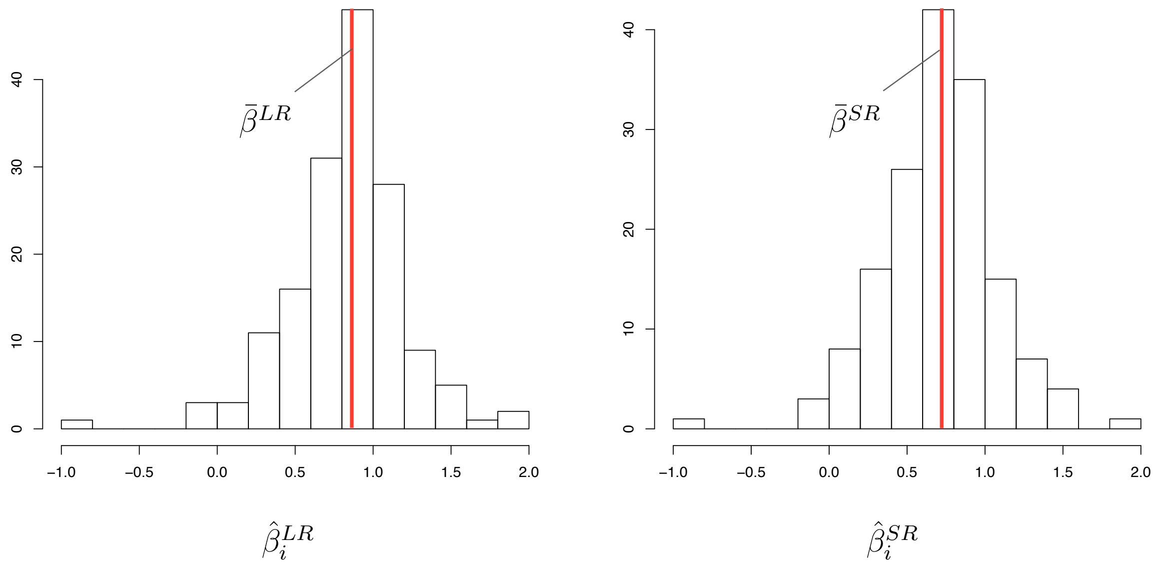

Figure 1: Distribution of country specific coefficient estimates: β_i^{LR} is estimated in equation (4) with the data in log-levels; β_i^{SR} is estimated in the error correction model in equation (6). $\bar{\beta}^{LR}$ and $\bar{\beta}^{SR}$ denote the average of the individual β_i^{LR} and β_i^{SR} coefficients, respectively.

and Pierucci and Ventura (2010)[3], respectively.

In addition to analyzing the results for any given economy, we can also compare the relative magnitude of long run and short run risk sharing across economies. The scatterplots in Figure 2 illustrate a positive relationship between the short and long run risk sharing coefficients. This pattern is independent of the period analyzed and is also present when the analysis is carried out for the sub-samples associated with the pre- and post-financial globalization environment. The positive correlation indicates that countries featuring higher short run betas, also tend to feature higher long run betas, or that countries experiencing less consumption smoothing in the short run also tend

[3]The formula for the mean adjustment lag used by Leibrecht and Scharler (2008) and by Pierucci and Ventura (2010) is based on the assumption of a homogeneous steady state, or $\hat{\beta}_i^{LR} = 1$. Because this assumption did not hold in their studies, we re-calculated the mean adjustment lags listed in their papers according the formula above.

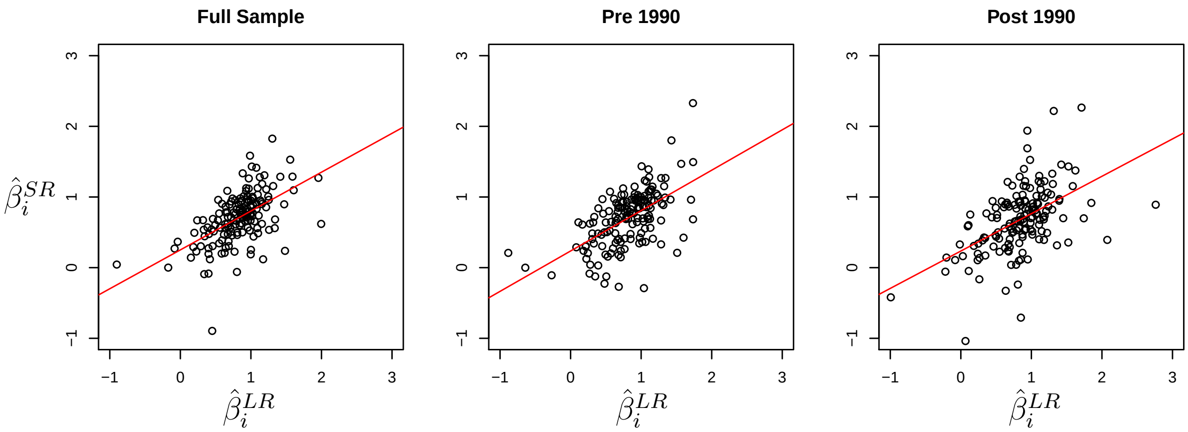

Figure 2: Relationship between short run and long run coefficients. The panels illustrate the relationship between coefficient estimates obtained from the entire sample, the pre-1990, and post-1990 sub-samples, respectively.

to do so in the long run. This is an interesting finding because insurance mechanisms may in principle work differently over different time horizons. For example, Asdrubali et al. (1996) argued that credit constraints can be more severe over longer time spans, and therefore the credit channel, which is key in buffering short run fluctuations, may be much less effective in absorbing permanent shocks. While it is quite likely that consumption insurance at different frequencies is governed by different mechanisms (see also Artis and Hoffmann, 2004; Becker and Hoffmann, 2006), our results suggest that there are no trade-offs of risk sharing at different frequencies. In the long run, countries generally don't make up for poor performance in the short run.

Table 4 contains the results of simple OLS regressions aimed at quantifying the relationship between $\hat{\beta}^{SR}$ and $\hat{\beta}^{LR}$. The column and row names reflect the regressands and regressors, respectively, with $pre90, post90, high$, and low/mid subscripts referring to the pre- and post-1990 subsample, the high, and the low and middle income countries, respectively. The results in Table 4 confirm a significant positive relationship between $\hat{\beta}_i^{SR}$ and $\hat{\beta}_i^{LR}$ with a coefficient smaller than one, indicating that an increase in $\hat{\beta}_i^{LR}$ tends to be accompanied by a somewhat smaller increase in $\hat{\beta}_i^{SR}$. The last two columns of Table 4, show that the relationship between short run and long run betas is stronger (steeper slope) for low and middle income countries.

For countries where $\hat{\beta}_i^{LR}$ exceeds $\hat{\beta}_i^{SR}$, the adjustment of consumption towards the steady state equilibrium is only partial at the time of an income shock. The remaining disequilibrium is eliminated in subsequent periods at the speed given by $\hat{\kappa}_i$. Correspondingly, there is an additional "short run" impact of income shocks on consumption beyond what is captured by $\hat{\beta}_i^{SR}$. Thus, even countries featuring statistically insignificant $\hat{\beta}_i^{SR}$'s, may experience an extended period of adjustments towards the long run equilibrium. Suppose, for example, that country i, currently in equilibrium, experiences a negative income shock, $\Delta y_{it} < 0$. Then, if $\beta_i^{SR} = 0$, consumption will be higher than the level predicted by the long-run relationship, and $\hat{\varepsilon}_{it+1}^{LR} > 0$. As $\kappa_i < 0$, this will bring about a decrease in consumption in the following period. Hence, even if $\hat{\beta}_i^{SR}$ is statistically insignificant or zero, the impact of a negative income shock may not be immediately and fully offset by an insurance mechanism, as one would conclude by only looking at $\hat{\beta}_i^{SR}$. Instead, whenever $\hat{\beta}_i^{SR} < \hat{\beta}_i^{LR}$, the adjustment will be spread over several periods. Therefore, in the presence of cointegration, the error correction dynamics implies that the magnitude of short run coefficients is insufficient to pin down the extent of consumption smoothing. As we have seen, risk sharing in steady state equilibrium is inextricably linked to its short run dynamics, and therefore we should not expect insurance mechanisms to work well in the long run, if they do not work well in the short run. In fact, the positive relationship between $\hat{\beta}^{SR}$ and $\hat{\beta}^{LR}$ demonstrates that countries performing poorly in the short run, relative to other countries, also have a poor performance over the long term.

One of the merits of working with such a heterogeneous dataset lies in the possibility of relating the variability of the β coefficients across countries to some key macroeconomic variables. Specifically, we characterize risk sharing in terms of 1) general development as captured by per capita GDP, 2) gross savings as a percentage of GDP, and 3) trade openness approximated by imports plus exports as a percentage of GDP. Gross savings as a percentage of GDP is a proxy for the maturity of a country's credit and financial system, and it signals the availability of a buffer that cushions the impact of shocks. Trade openness measures the general attitude towards exchanging resources (and therefore also risk) with other countries, which is considered a relevant risk sharing channel in the literature.

Table 5 presents coefficient estimates in cross-sectional regressions of $\hat{\beta}_i^{SR}$ and $\hat{\beta}_i^{LR}$ on time averages of various macroeconomic indicators in the form of equations (8), where the $x_{\cdot i}$ regressors include GDP per capita, the percentage of gross savings over GDP, the degree of openness of country i, and the interactions of these variables. On their own, greater prosperity, savings, and openness to the rest of the world appear to be associated with a significant reduction in $\hat{\beta}^{SR}$ or an increase in short run risk sharing. While the signs of both short and long run estimates are the same, the latter are not statistically significant at conventional levels. However, the last two columns of Table 5 suggest that when additional factors are taken into account, GDP per capita does not contribute to a larger degree of international risk sharing. In fact, the coefficients of GDP per capita are positive (and significant in the short run) implying that it is inversely related to risk sharing. On the other hand, the significant negative coefficient of the interaction term between gross savings as a fraction of GDP and GDP per capita shows that the size of the savings buffer in the economy is key in fostering risk sharing. This notion is reinforced by the negative coefficient of the gross savings as a fraction of GDP variable. What drives risk sharing up, therefore, is not so much a general level of development, but rather its combination with the depth of domestic credit and financial markets.

4 Conclusion

Many recent studies have estimated international risk sharing under the assumptions of homogeneous economies. Our main contribution to the literature lies in accounting for the inevitable heterogeneity of a dataset containing information on multiple countries. Even datasets that are smaller than the one in our study usually contain economies that differ widely with respect to a number of relevant features. Such heterogeneity, if not properly accounted for, might bias the coefficient estimates and produce incorrect conclusions about the extent of risk sharing. We properly account for country heterogeneity and estimate risk sharing in a large set of diverse economies using modern econometric tools. Although our results confirm that risk sharing is partial at best, we find a large variation in the level of risk sharing across countries.

Furthermore, we embed the analysis in a dynamic framework and show that long term and short term risk sharing are inextricably linked. We find that countries performing poorly in the short run, relative to other countries, also tend to have a poor performance over the long run. This bears a remarkable policy implication: taking advantage of short term insurance opportunities has long term benefits. Finally, the large data panel in our study allows us to identify macroeconomic factors fostering better consumption insurance. We observe that the extent of development, per se, is not enough to secure a higher level of consumption smoothing unless economic development is accompanied by a well functioning credit market. Consequently, policymakers should implement measures that expand and facilitate access to credit markets.

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Table 2: Comparison of country-specific coefficient estimates

id	country	β^{LR}	β^{SR}	κ	id	country	β^{LR}	β^{SR}	κ
1	AFG	0.87*	0.95*	-0.20 *	41	DOM	0.90*	1.01*	-0.21 *
2	AGO	2.00*	0.62	-0.58 *	42	DZA	1.48*	0.24	-0.21
3	ALB	0.18*	0.29*	-0.22 *	43	ECU	0.71*	0.57*	-0.58 *
4	ARG	0.97*	1.26*	-0.19 *	44	EGY	0.68*	0.37	-0.33 *
5	ATG	0.99*	1.58*	-0.70 *	45	ESP	0.89*	0.80*	-0.32 *
6	AUS	0.42*	0.12	-0.10	46	ETH	1.10*	1.03*	-0.43 *
7	AUT	1.11*	0.74*	-0.35 *	47	FIN	0.81*	0.46*	-0.27 *
8	BDI	0.86*	0.79*	-0.52 *	48	FJI	0.64*	0.57*	-0.37 *
9	BEL	0.98*	0.53*	-0.35 *	49	FRA	1.02*	0.73*	-0.24 *
10	BEN	0.80*	0.66*	-0.22 *	50	FSM	1.03*	0.98*	-0.81 *
11	BFA	1.96*	1.27*	-0.52 *	51	GAB	0.34*	-0.09	-0.56 *
12	BGD	1.61*	1.10*	-0.62 *	52	GBR	1.13*	0.95*	-0.55 *
13	BGR	1.01*	0.85*	-0.39 *	53	GER	1.16*	0.62*	-0.28 *
14	BHR	0.64*	0.78*	-0.42 *	54	GHA	1.00*	0.88*	-0.65 *
15	BHS	1.42*	1.29*	-0.36 *	55	GIN	1.56*	1.53*	-0.33 *
16	BLZ	1.22*	1.11*	-0.56 *	56	GMB	0.85*	0.92*	-0.35 *
17	BMU	1.47*	0.90*	-0.38 *	57	GNB	0.99*	0.74*	-0.50 *
18	BOL	0.64*	0.86*	-0.46 *	58	GNQ	0.67*	0.72*	-0.31 *
19	BRA	0.83*	0.90*	-0.46 *	59	GRC	0.20	0.49*	-0.15 *
20	BRB	1.59*	1.29*	-0.31 *	60	GRD	0.46*	0.69*	-0.53 *
21	BRN	-0.90*	0.04	-0.57 *	61	GTM	0.86*	0.75*	-0.26 *
22	BTN	0.56*	0.57*	-0.41 *	62	GUY	1.15*	0.93*	-0.26 *
23	BWA	0.40*	0.19*	-0.34 *	63	HKG	1.22*	0.85*	-0.43 *
24	CAF	0.92*	0.83*	-0.45 *	64	HND	1.00*	0.19	-0.54 *
25	CAN	0.41*	0.48*	-0.18 *	65	HTI	0.97*	1.12*	-0.46 *
26	CHE	0.23	0.23*	-0.32 *	66	HUN	1.02*	1.00*	-0.20 *
27	CHL	0.96*	0.79*	-0.22	67	IDN	1.09*	0.56*	-0.27 *
28	CHN2	1.00*	1.01*	-0.42 *	68	IND	0.95*	0.71*	-0.58 *
29	CIV	0.76*	0.78*	-0.48 *	69	IRL	0.64*	0.62*	-0.48 *
30	CMR	0.97*	0.79*	-0.40 *	70	IRN	0.59*	0.39*	-0.27 *
31	COG	0.46*	0.30*	-0.11	71	IRQ	-0.08	0.27*	-0.86 *
32	COL	0.92*	0.75*	-0.40 *	72	ISL	1.25*	1.01*	-0.60 *
33	COM	0.40*	0.27	-0.38 *	73	ISR	1.08*	0.83*	-0.39 *
34	CPV	0.99*	0.67*	-0.13	74	ITA	1.00*	0.76*	-0.53 *
35	CRI	0.93*	1.13*	-0.18	75	JAM	0.88*	0.80*	-0.43 *
36	CUB	1.16*	1.18*	-0.34 *	76	JOR	1.34*	0.68*	-0.30 *
37	CYP	0.87*	0.72*	-0.48 *	77	JPN	0.86*	0.63*	-0.27 *
38	DJI	1.25*	0.96*	-0.72 *	78	KEN	1.08*	1.41*	-0.43 *
39	DMA	0.62*	0.47*	-0.70 *	79	KHM	0.94*	0.97*	-0.57 *
40	DNK	0.53*	0.67*	-0.32 *	80	KIR	0.67*	0.47*	-0.30 *

Notes: β^{LR} denotes the CCE estimate with the data in log-levels, as in equation (4). β^{SR} denotes the estimate in Error-Correction Model, or equation (6). κ denotes the speed-of-adjustment estimate in the Error-Correction Model. Robust standard errors (HAC) are used for inference. Statistical significance at the 5% level or lower is denoted by *.

Table 3: Comparison of country-specific coefficient estimates (continued)

id	country	β^{LR}	β^{SR}	κ	id	country	β^{LR}	β^{SR}	κ
81	KNA	0.80*	-0.06	-0.51 *	121	PRT	0.81*	0.50*	-0.49 *
82	KOR	0.84*	0.78*	-0.19 *	122	PRY	0.77*	0.23	-0.30 *
83	LAO	0.85*	0.98*	-0.39 *	123	ROM	0.50*	0.61*	-0.39 *
84	LBN	0.84*	0.84*	-0.81 *	124	RWA	0.59*	0.20*	-0.11 *
85	LBR	1.15*	0.90*	-0.28 *	125	SDN	1.01*	1.43*	-0.34 *
86	LCA	0.70*	0.89*	-0.52 *	126	SEN	0.88*	0.50*	-0.21 *
87	LKA	1.10*	0.49*	-0.62 *	127	SGP	0.73*	0.46*	-0.13
88	LSO	1.05*	0.66*	-0.35 *	128	SLB	1.04	0.44*	-0.17 *
89	LUX	0.69*	0.31*	-0.26 *	129	SLE	0.67*	0.79*	-0.23 *
90	MAC	1.00*	0.25*	-0.11 *	130	SLV	1.32*	1.16*	-0.55 *
91	MAR	0.34	0.44*	-0.50 *	131	SOM	0.95*	1.01*	-0.48 *
92	MDG	1.17*	0.12	-0.10	132	STP	1.19*	1.31*	-0.63 *
93	MDV	0.96*	0.66*	-0.24 *	133	SUR	1.30*	1.83 *	-0.68 *
94	MEX	0.80*	0.83*	-0.19 *	134	SWE	0.67*	0.58 *	-0.32 *
95	MHL	1.34*	0.56*	-0.40 *	135	SWZ	0.40*	-0.08	-0.66 *
96	MLI	-0.17	0.00	-0.44 *	136	SYC	1.08*	0.89*	-0.23
97	MLT	0.75*	0.67*	-0.13	137	SYR	0.78*	0.96*	-0.20 *
98	MNG	0.67*	1.09*	-0.52 *	138	TCD	0.47*	0.51*	-0.19
99	MOZ	0.89*	0.70*	-0.44	139	TGO	0.33 *	0.54*	-0.83 *
100	MRT	0.80*	0.83*	-0.49 *	140	THA	0.78*	0.61*	-0.70 *
101	MUS	0.84*	0.60*	-0.47 *	141	TON	0.92*	0.65*	-0.56 *
102	MWI	0.44*	0.54*	-0.86 *	142	TTO	0.93*	0.81*	-0.77 *
103	MYS	0.54*	0.96*	-0.26 *	143	TUN	0.64*	0.28*	-0.53 *
104	NAM	1.03*	0.88*	-0.85 *	144	TUR	0.72*	0.94*	-0.38 *
105	NER	0.32*	0.67*	-0.62 *	145	TWN	1.12*	0.63*	-0.37 *
106	NGA	1.13*	1.28*	-0.45 *	146	TZA	-0.04	0.37*	-0.33 *
107	NIC	0.72*	0.51*	-0.36 *	147	UGA	0.97*	0.94*	-0.66 *
108	NLD	0.72*	0.71*	-0.18 *	148	URY	1.00*	0.95*	-0.05
109	NOR	0.38*	0.57*	-0.26 *	149	USA	0.91*	0.74*	-0.21
110	NPL	1.10*	1.13*	-0.48 *	150	VCT	0.74*	0.98*	-0.36 *
111	NZL	0.95*	0.79*	-0.36 *	151	VEN	1.02*	0.77*	-0.35
112	OMN	1.26*	0.53	-0.57 *	152	VNM	0.55*	0.77*	-0.14
113	PAK	0.59*	0.90*	-0.49 *	153	VUT	0.78*	0.76*	-0.32 *
114	PAN	0.15	0.14	-0.65 *	154	WSM	0.95*	0.92*	-0.67 *
115	PER	0.95*	0.93*	-0.36 *	155	ZAF	0.67*	0.62*	-0.38 *
116	PHL	0.29*	0.30*	-0.21 *	156	ZAR	0.63*	0.21	-0.55 *
117	PLW	0.45	-0.89 *	-0.12	157	ZMB	0.88*	1.34*	-0.33 *
118	PNG	0.96	0.57	-0.08	158	ZWE	0.24	0.67*	-0.52 *
119	POL	0.93*	1.08*	-0.40 *					
120	PRI	0.55*	0.34*	0.01		OECD	0.80	0.68	-0.31

Notes: See also the notes in Table 2. Coefficient values for OECD have been computed as mean group estimates. Robust standard errors (HAC) are used for inference. Statistical significance at the 5% level or lower is denoted by *.

Table 4: Short run vs. long run risk sharing

REGRESSORS	REGRESSANDS β^{SR}	β^{SR}_{pre}	β^{SR}_{post}	β^{SR}_{high}	$\beta^{SR}_{low/mid}$
β^{LR}	0.550***				
	(0.0753)				
β^{LR}_{pre90}		0.572***			
		(0.0812)			
β^{LR}_{post90}			0.529***		
			(0.0903)		
β^{LR}_{high}				0.471***	
				(0.0845)	
$\beta^{LR}_{low/mid}$					0.583***
					(0.103)
Constant	0.251***	0.236***	0.235***	0.270***	0.239***
	(0.0631)	(0.0648)	(0.0752)	(0.0704)	(0.0855)
R-squared	0.316	0.334	0.260	0.493	0.289

Notes: This table presents the basic results of simple regressions relating short run smoothing coefficients to long run ones. The *pre*90, *post*90, *high*, and *low/mid* subscripts refer to the pre- and post-globalization subsample, the high, and the low and middle income countries, respectively. Robust standard errors in parentheses. The marginal level of significance at the 1%, 5%, and 10% is denoted by ***, **, *, respectively.

Table 5: Drivers of risk sharing

REGRESSORS	REGRESSANDS β^{SR}	β^{LR}	β^{SR}	β^{LR}	β^{SR}	β^{LR}	β^{SR}	β^{LR}
GDP	-3.81e-06*	-1.79e-06					.00001***	.00002*
	(1.98e-06)	(2.53e-06)					(4.22e-06)	(.00001)
saving			-.0107***	-.00681			-.0062**	.0007
			(0.0026)	.(0046)			(0.0031)	(0.0035)
openness					-.0011**	-.00031	-0.00008	0.00061
					(0.0005)	(0.0005)	(0.00054)	(0.00064)
GDP*saving							-5.99e-07***	-1.10e-06*
							(1.60e-07)	5.65e-07
Constant	.737***	.845***	.921***	.954***	.800***	.849***	.845***	.819***
	(0.0631)	(0.0648)	(0.0752)	(0.0704)	(.0496)	(0.0508)	(.0698)	(0.0728)
R-squared	0.0155	0.0032	0.083	0.029	0.0203	0.0015	0.109	0.100

Notes: This table presents the basic results of simple regressions relating short and long run smoothing coefficients to a set of macro variables. Robust standard errors in parentheses. The marginal level of significance at the 1%, 5%, and 10% is denoted by ***, **, *, respectively.

Appendix: List of countries by sub-group[4]

High Income: Australia, Austria, Bahamas, Bahrain, Barbados, Belgium, Bermuda, Brunei, Canada, Cyprus, Denmark, Equatorial Guinea, Finland, France, Germany, Greece, Hong Kong, Hungary, Iceland, Ireland, Israel, Italy, Japan, Republic of Korea, Luxembourg, Macao, Malta, Netherlands, New Zealand, Norway, Oman, Poland, Portugal, Puerto Rico, Singapore, Spain, St. Kitts & Nevis, Sweden, Switzerland, Trinidad & Tobago, United Kingdom, United States.

Upper Mid-Income: Algeria, Angola, Antigua & Barbuda, Argentina, Botswana, Brazil, Bulgaria, Chile, China, Colombia, Costa Rica, Cuba, Dominica, Dominican Republic, Ecuador, Gabon, Grenada, Iran, Jamaica, Jordan, Lebanon, Malaysia, Maldives, Mauritius, Mexico, Namibia, Palau, Panama, Peru, Romania, Seychelles, South Africa, St. Lucia, St.Vincent & Grenadines, Suriname, Thailand, Tunisia, Turkey, Uruguay, Venezuela.

Lower Mid-Income: Albania, Belize, Bhutan, Bolivia, Cameroon, Cape Verde, Republic of Congo, Cote d'Ivoire, Djibouti, Egypt, El Salvador, Fiji, Ghana, Guatemala, Guyana, Honduras, India, Indonesia, Iraq, Kiribati, Lesotho, Marshall Islands, Fed. Sts. Micronesia, Mongolia, Morocco, Nicaragua, Nigeria, Pakistan, Papua New Guinea, Paraguay, Philippines, Samoa, Sao Tome & Principe, Senegal, Solomon Islands, Sri Lanka, Sudan, Swaziland, Tonga, Vanuatu, Vietnam, Zambia.

Low Income: Afghanistan, Bangladesh, Benin, Burkina Faso, Burundi, Cambodia, Central African Republic, Chad, Comoros, Dem. Rep. Congo, Ethiopia, Gambia, The, Guinea, Guinea-Bissau, Haiti, Kenya, Liberia, Madagascar, Malawi, Mali, Mauritania, Mozambique, Nepal, Niger, Rwanda, Sierra Leone, Somalia, Tanzania, Togo, Uganda, Zimbabwe.

OECD: Australia, Austria, Belgium, Canada, Chile, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Republic of Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, Turkey, United Kingdom, United States.

OECD-AH: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom, United States.

Developed: Australia, Austria, Belgium, Canada, Cyprus, Denmark, Finland, France, Germany, Greece, Hungary, Israel, Ireland, Italy, Japan, Luxembourg, Malta,

[4]Our conclusions are robust to the presence of small countries in the sample. The results do not materially change when countries with population below 1 million are excluded from the analysis.

Netherlands, New Zealand, Norway, Poland, Portugal, Romania, Spain, Sweden, Switzerland, United Kingdom, United States.

Developing: Algeria, Angola, Argentina, Bahrain, Bangladesh, Barbados, Benin, Bermuda, Bolivia, Botswana, Brazil, Burkina Faso, Burundi, Cameroon, Cape Verde, Central African Republic, Chad, Chile, China, Colombia, Comoros, Dem. Rep. Congo, Republic of Congo, Costa Rica, Cote d'Ivoire, Cuba, Djibouti, Dominican Republic, Ecuador, Egypt, El Salvador, Equatorial Guinea, Ethiopia, Gabon, The Gambia, Ghana, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, Hong Kong, India, Indonesia, Iran, Iraq, Israel, Jamaica, Jordan, Kenya, Republic of Korea, Lebanon, Lesotho, Liberia, Madagascar, Malawi, Malaysia, Mali, Mauritania, Mauritius, Mexico, Morocco, Mozambique, Namibia, Nepal, Nicaragua, Niger, Nigeria, Oman, Pakistan, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Rwanda, Sao Tome and Principe, Senegal, Sierra Leone, Singapore, Somalia, South Africa, Sri Lanka, Sudan, Taiwan, Tanzania, Thailand, Togo, Trinidad & Tobago, Tunisia, Turkey, Uganda, Uruguay, Venezuela, Vietnam, Zambia, Zimbabwe.

EU: Austria, Belgium, Bulgaria, Cyprus, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Luxembourg, Malta, Netherlands, Poland, Portugal, Romania, Spain, Sweden, United Kingdom.